

Some New Estimates of the Navy's Indirect Manning Costs

Henry L. Eskew

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Introduction and summary

How does a change in the manning of ships and squadrons at sea affect the Navy's shore-based manning? This question, while hardly new, has arisen recently in several different contexts. One involves cost-effectiveness analyses of *arsenal* ships—which require relatively small crews—as alternatives to traditional surface combatants.

The question also arises in connection with tradeoffs between support ships that are Navy owned and manned and support ships that are leased annually and manned by civilian crews. A third setting involves the use of aggregate models to project long-term fiscal requirements generated by alternative force structures that vary in both size and composition. Such models must necessarily capture the totality of personnel costs, and thus require a mechanism for quantifying the relationship between manning ashore and at sea.

From what we have just said, it is clear that the interest here has to do largely with cost analysis. In particular, the focus is on what are loosely referred to as *indirect* costs. In the past, for reasons that are not entirely apparent, the emphasis placed on identification and estimation of indirect costs has been at best sporadic. In many analyses, even when costs are an important input, ignoring indirect manning costs may be entirely reasonable. A good example is a cost-effectiveness analysis of alternative attack aircraft, where the acquisition costs and performance capabilities of the aircraft vary widely, but their direct—and hence indirect—manning requirements are not measurably different. However, in the types of settings we have just described, ignoring indirect costs would clearly distort the analysis.

The purpose of this paper is to provide a set of empirical estimates of the response of ashore manning to changes in manning of ships and squadrons—hereafter called *afloat* manning—based on the most recent time-series information available. Over the past six or seven years, the drawdowns in budgets, force structure, and manning have

been substantial. Inclusion of that experience in the database from which cost-estimating relationships are developed is essential to the validity of the relationships for use in assessing the cost consequences of decisions presently or soon to be at hand.

The analytical construct adopted here is a model that posits delayed adjustment of ashore manning to changes in afloat manning. This so-called *partial adjustment* model is tractable statistically but must nevertheless be treated with a degree of caution. We employed a time series that spans fiscal years 1980 through 1996 and took data from the historical and current releases of the Future Years Defense Program (FYDP). Our principal findings may be summarized as follows:

- When afloat manning is changed by N officers, the full (as opposed to initial) response ashore is a change of approximately 90 percent as many officers.
- When afloat manning is changed by N enlistees, the full (as opposed to initial) response ashore is a change of approximately N enlistees.
- The adjustment process is substantially faster for officers than for enlistees.
- Officers ashore exhibit a large and statistically significant fixed component—in the neighborhood of 30,000—but there is no statistical evidence of a fixed component of enlistees.

The first two findings have important implications for cost analysis. The resultant indirect personnel *multipliers*—0.9 for officers and 1.0 for enlistees—are substantially higher than any that are known to be in use now or in the past. The third and fourth findings mainly provide statistical confirmation and definition of certain perceptions that are not likely to be in dispute.

The model

Without a doubt, a relationship exists between the number of people afloat and the number ashore. The Navy's management of human resources is built around a complex system of *rotations*. One objective of the system is to ensure an equitable balance between time spent at sea and ashore. Another is to effect the rotations necessary for the orderly career progression of both officers and enlistees. This also involves rotation between sea and shore duty.

When the numbers and types of operating forces change from one fiscal year to the next, it is easier for the Navy to make the necessary adjustments in manning of ships and squadrons than to adjust the remainder—and hence the total size—of the force. The size of the total force is limited by a host of institutional arrangements, not the least of which is the role of Congress in authorizing personnel endstrengths. In addition, there are:

- Contractual agreements affecting length of service
- Restrictions on the availability of funds for accessions and separations
- Lead times required for the inputs and outputs of the various pipelines to reach equilibrium.

All of this suggests that any model of the process must take explicit account of the lag, or delay, in the adjustment of shore-based manning to changes affecting ships and squadrons. One such formulation is the partial adjustment model, which has been applied widely in the econometric literature—see [1], for example—and was used in a previous analysis of the same question addressed in this paper [2]. The latter reference also reviews earlier studies of the adjustment of resources ashore.

As applied in the present context, the model posits that there is a desired level of ashore manning in year t , denoted by S_t^* , which is a linear function of afloat manning in the same year, F_t , i.e.,

$$S_t^* = \alpha + \beta F_t .$$

However, because of the institutional constraints discussed above, the *observed* adjustment, $(S_t - S_{t-1})$ in any one year consists of some fixed fraction (λ) of the desired adjustment plus a random error, u_t :

$$(S_t - S_{t-1}) = \lambda (S_t^* - S_{t-1}) + u_t ,$$

where $0 < \lambda < 1$. The closer λ is to 1, the faster the adjustment. Combining these two equations gives

$$S_t = \alpha\lambda + \beta\lambda F_t + (1 - \lambda) S_{t-1} + u_t .$$

The structural parameters that reflect the speed and magnitude of the full adjustment process appear in this equation, although in non-linear form. The equation may, however, be rewritten in the following notation:

$$S_t = \gamma_0 + \gamma_1 F_t + \gamma_2 S_{t-1} + u_t .$$

Observing that $\gamma_0 = \alpha\lambda$, $\gamma_1 = \beta\lambda$, and $\gamma_2 = (1 - \lambda)$, it follows that

$$\alpha = \gamma_0 / (1 - \gamma_2), \quad \beta = \gamma_1 / (1 - \gamma_2), \quad \text{and} \quad \lambda = (1 - \gamma_2) .$$

Thus if we can obtain estimates of the γ 's that have desirable statistical properties, we can use those estimates to derive estimates of the structural parameters. The two sets of estimates will then have the same properties. We will explore these matters in the section on estimation issues and results.

The data

We can measure naval manning in several different ways, but the interpretation of any given set of manning data is not always clear. One measurement concept focuses on requirements, which tend to be denominated in billets. Some analysts like to think of billets as being akin to *chairs* provided for people to occupy. In some cases, there are more chairs than people; in others there are less.

If our objective is to measure numbers of people actually employed, we encounter the dichotomy between person-years and endstrength. The former, a flow concept, is more closely correlated with the magnitude of annual pay and allowances. It is also the more elusive from a measurement perspective. Endstrength is a stock measure. It represents an inventory count at the end of each fiscal year, and is the more operational of the two. On the premise that it is people, not billets, that generate costs, we have adopted endstrength as the measurement concept for this analysis. From the point of view of quantifying adjustment processes, that concept has the added advantage of allowing the maximum amount of time for the processes to reach their outcomes in each fiscal year observed. (The use of 12-month inventory averages would hide a portion of the adjustment.)

FYDPs are organized by program element (PE) within 11 major program areas. Those documents report endstrength manning for each PE in the Navy. For the program elements that represent the fleet's ships and squadrons, virtually all active-duty manning—the focus of this study—appears in programs 1 and 2, Strategic Forces and General Purpose Forces, respectively.¹ The (separate) sums of officers and enlistees associated with ships and squadrons for each fiscal year

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1. An exception is the aircraft carrier, presently CV-67, assigned to the reserve forces. Substantial numbers of active duty personnel appear against that PE and were included in this study's definition of afloat manning.

constitute our definition of afloat manning. We obtained our measures of ashore manning by subtracting the afloat sums from total Navy manning reported in the same documents. Data for fiscal 1980 through 1992 came from the most recent release of the historical FYDP [3]. We used the February 1994 release [4] to obtain 1993 data and the February 1995 issue [5] to complete the time series.² Those data are reprinted in appendix A.

Figure 1 is a plot of officers afloat (left panel) and ashore (right panel) over time. The greater uniformity in the former bears out the earlier discussion of the greater ease with which the Navy can adjust to changes at sea than ashore. The fact that manning afloat peaks two years earlier than manning ashore (1988 as compared with 1990) is further evidence of the delay in adjustment. The statistical results presented later will further illuminate and quantify the lag in the adjustment process.

The plots of enlistee data in figure 2 display essentially the same patterns as figure 1, although the difference in uniformity is perhaps not as great. The two-year lag in peak manning ashore and afloat is still present although, in this case, the peaks occur in 1987 and 1989.

In figure 3, ashore and afloat manning are plotted against one another, with officers in the left panel and enlistees in the right. Those plots are clearly consistent with the hypothesized linear relationship between the two pairs of variables, as qualified by the preceding discussion and figures. The task of the statistical analysis that follows is to estimate and test the parameters in those relationships.

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2. Earlier studies of this issue have used time series that included much of the 1970s. In the present case, the historical FYDP extended back only through 1980. More importantly, however, there is reason to believe that relatively little is lost by omitting those years, and in fact their exclusion may be an advantage. First, a good portion of that decade was characterized by relatively flat manning ashore and at sea. Hence that information contributes little to the measurement of the response of the former to changes in the latter. In addition, the first part of the 1970s experienced the drawdown from the Viet Nam conflict coupled with arrangements that existed before the all-volunteer force was introduced. It is doubtful that the same statistical structure that applies in the 1980s and 1990s also applied in that period.

Figure 1. Officer manning afloat and ashore over time

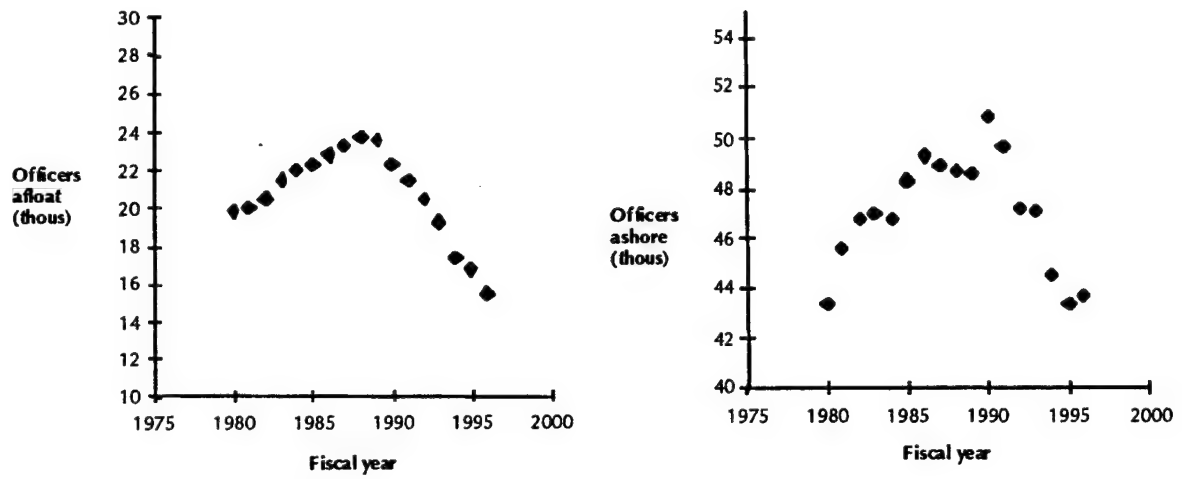


Figure 2. Enlisted manning afloat and ashore over time

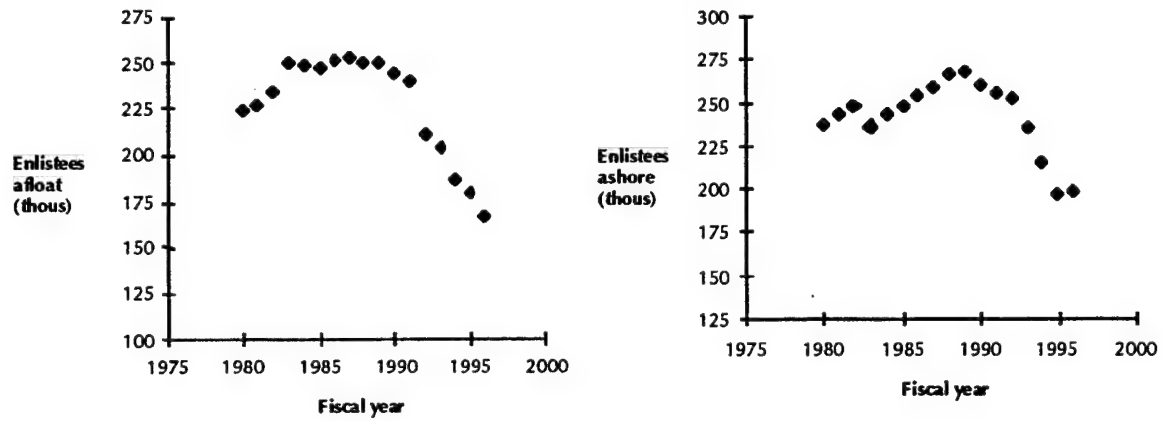
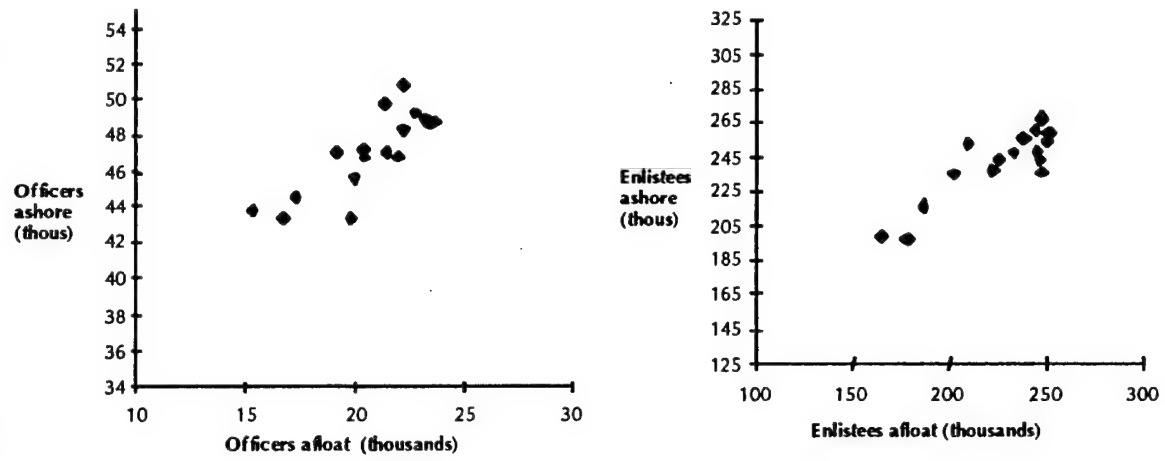


Figure 3. Ashore versus afloat manning, officers and enlistees



Estimation issues and results

A useful place to start is with simple regressions of ashore manning on afloat manning for each set of data. We have made the *a priori* case—supported by the preceding scatter plots—that such regressions would mask the more complicated process underlying the data. The question now is whether we can find further evidence of that assumption through formal statistical analysis. Table 1 lists the regression results.

Table 1. Simple regressions of ashore manning on afloat manning

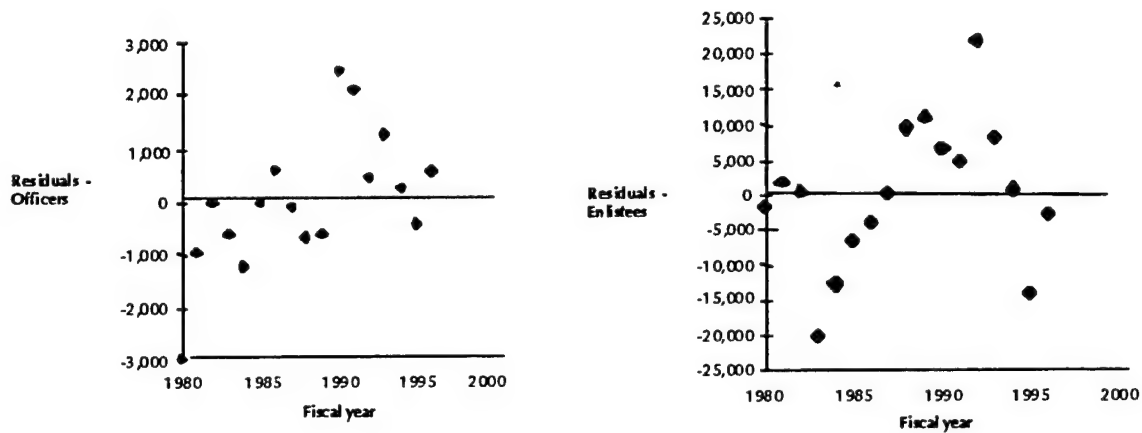
Data set	Intercept	Slope	S.E.E.	Dep. mean	Ind. mean	\bar{R}^2	D.W.
Officers	30,625 (10.26)	0.796 (5.56)	1,397	47,109	20,698	0.651	0.72
Enlistees	93,157 (4.06)	0.656 (6.54)	11,155	242,234	227,126	0.723	0.99

On the surface, these results appear quite reasonable. Each of the parameter estimates is intuitively plausible and highly significant as judged by the *t* ratios in parentheses, which are well above the conventional rule-of-thumb value of 2.0. In addition, the standard errors of estimate (S.E.E.) are relatively small—less than 5 percent of the mean value of the dependent variable in each case—and the \bar{R}^2 statistics, although not especially large, are certainly adequate. However, the very low Durbin-Watson (D.W.) values, indicating the presence of positive autocorrelation among the errors, constitute a problem—as is often the case with time series data.³ (An ideal value for the D.W.

3. At the risk of oversimplification, the Durbin-Watson statistic is used to test the hypothesis that $\rho = 0$ in the model, $u_t = \rho u_{t-1} + v_t$. Rejection of that hypothesis in favor of $\rho > 0$, which is indicated here by reference to standard tables of D.W. values, establishes the presence of positive autocorrelation.

statistic is roughly 2.0) A lesser consequence of this is that the t ratios can no longer be given their usual interpretation. A considerably more important implication for our purposes is found in the patterns of residuals from the regressions. The residuals constitute estimates of the unobservable random errors and are the basis from which the D.W. statistics are computed. Those values, computed as $(S_{act} - S_{est})$, are plotted in figure 4 as a function of time.

Figure 4. Residuals from the simple regressions



For both officers and enlistees, the residuals are mostly negative in the first several years of the time series—the period in which afloat manning was increasing—and tend to be positive during the years when manning at sea was declining. In other words, manning ashore during the growth period fell short of the levels predicted by the regression, whereas the opposite occurred during the period of decline. This is still further evidence of the lag in the adjustment process. We thus shift attention to the partial adjustment model.

As noted earlier, that model may be written as

$$S_t = \gamma_0 + \gamma_1 F_t + \gamma_2 S_{t-1} + u_t ,$$

where the structural parameters of interest— α , β , and λ —are imbedded but recoverable. We now consider the issues associated with statistical estimation of this model.⁴

The first issue concerns the presence of a lagged value of the dependent variable on the right side of the equation. The consequence of this is that the least-squares estimates of the parameters will exhibit bias in small samples, but the bias will tend to vanish as the sample size becomes large.⁵ Thus the estimates have the desirable property of *consistency*.⁶ There is a tendency to think that because consistency is a large-sample property, its presence or absence may be irrelevant in samples as small as those here. However, evidence to the contrary is found in the results of Monte Carlo experiments presented in [6, pp. 739–742], where consistent estimators were seen to perform considerably better than inconsistent ones even in samples roughly the same size as these.

A second and equally important issue is that the consistency of the least-squares estimates (in the presence of a lagged dependent variable) is crucially dependent on another condition: the absence of autocorrelated errors. The next order of business then is to estimate the partial adjustment model and test for autocorrelation. We first consider the results in table 2.

We note the improvement over the earlier results in the S.E.E. and \bar{R}^2 values, each of which is adjusted for degrees of freedom. Thus the partial adjustment model constitutes a better fit to the data. Before commenting on the D.W. statistics, we see from the plots in figure 5 that there are no discernible patterns in the residuals. This constitutes a kind of visual confirmation of the absence of autocorrelation.

4. These issues are discussed in virtually all econometrics texts. See, for example, [6, pp. 737–744].

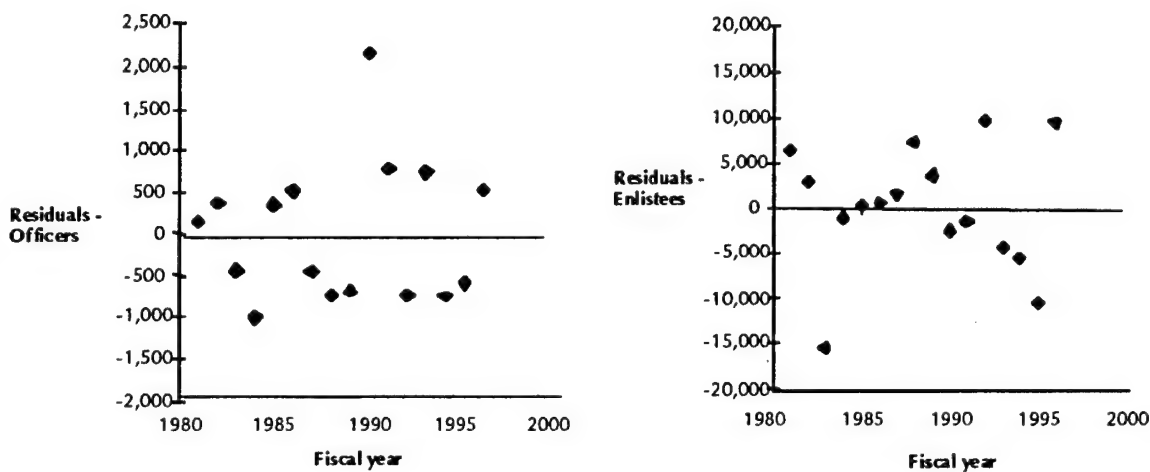
5. If the random variable b is an estimator of the parameter β , and if the expected value of b is not equal to β , then b is said to be a biased estimator of β .

6. Consistency is a moderately technical concept; this is an extremely intuitive definition.

Table 2. Estimates of the partial adjustment model

Data set	Est. of $\gamma_0 = \alpha\lambda$	Est. of $\gamma_1 = \beta\lambda$	Est. of $\gamma_2 = 1 - \lambda$	S.E.E.	Dep. mean	Ind. mean	\bar{R}^2	D.W.
Officers	15,111 (3.06)	0.511 (4.50)	0.457 (3.58)	862	47,347	20,757	0.849	1.77
Enlistees	5,264 (0.19)	0.358 (3.41)	0.636 (3.93)	8,089	242,564	227,408	0.863	2.08

Figure 5. Residuals from the partial adjustment model



It turns out that when there is a lagged dependent variable in the equation, the Durbin-Watson test is not likely to be valid.⁷ When such is the case, we need an alternative test. The one we used here was recommended in [8, p. 454]. The outcomes of the tests were highly conclusive in support of the absence of autocorrelated errors.⁸

7. This matter is examined in [7].

8. The test consists of first regressing the residuals on F_t , S_{t-1} , and the residuals lagged one year. Then the significance of the resultant set of regression coefficients is tested by means of standard F test. The computed F values were very close to zero, indicating almost total absence of statistical significance, and hence absence of autocorrelation.

The previous discussion indicates that the estimates of the structural parameters derived from the regression results in table 2 will be consistent. We list those estimates in table 3 below, along with estimates of the elasticity of the full response of ashore manning to changes in afloat manning, computed at the sample means of the variables.⁹

Table 3. Estimates of the structural parameters and elasticities

Data set	α	β	λ	Elasticity
Officers	27,829	0.941	0.543	0.413
Enlistees	14,461	0.984	0.364	0.923

One piece of information is still missing from this picture: measures of the statistical significance of the estimates of the structural parameters. Each of the parameter estimates in table 2 had highly significant *t* ratios, except for the intercept term in the equation for enlistees. But those estimates pertain to parameters that represent nonlinear functions of α , β , and λ . Therefore, quantification of their sampling errors and significance measures by standard methods is not easily achieved. However, a straightforward approach that will shed considerable light on the matter is to specify the model in its original form,

$$S_t = \alpha\lambda + \beta\lambda F_t + (1 - \lambda) S_{t-1} + u_t ,$$

and then estimate α , β , and λ directly by nonlinear least squares. The asymptotic properties of those estimators—consistency, normality, and (in special cases) efficiency—are well known, and the asymptotic standard errors can serve to at least approximate the significance of the estimates.¹⁰ Those results are in table 4.

9. That computation is performed by multiplying the estimated value of β by the ratio of the mean of afloat manning to the mean of ashore manning.

10. Properties of the nonlinear least-squares estimators are summarized in [8, pp. 335–340].

Table 4. Nonlinear least-squares estimates

	Officers		Enlistees	
	Param. est.	Asymptotic standard error	Param. est.	Asymptotic standard error
α	27,825	3,906	14,475	71,846
β	0.941	0.187	0.985	0.305
λ	0.543	0.127	0.364	0.162

The fact that these estimates of α , β , and λ are virtually identical to those in table 3 is no coincidence. Because the least-squares criterion was applied to generate the estimates in tables 2 and 4, and because the estimates in table 3 were uniquely recoverable from those in table 2, the two sets must necessarily conform. We note that, with the exception of α in the enlistee equation, the parameter estimates are several times as large as their asymptotic standard errors. This should settle any remaining questions about the significance of those estimates, although we must conclude that there is no statistical evidence of a fixed component of enlistees ashore.¹¹ We also infer from the estimates of λ that the adjustment of officers ashore is half again as fast as that of enlistees.

11. We do not rule out the possibility that a fixed component of enlistees would emerge if the drawdown in force structure were extended well beyond the range observed in this study.

Sensitivity analysis

Two areas of sensitivity analysis suggest themselves. The first concerns the span of years included in the database, and the second has to do with the definition of afloat manning. Statistical results supporting the analysis are in appendix B. The following is a brief summary of its motivation and outcome.

We took manning data for fiscal 1995 and 1996 from the FYDP that was released in February 1995 and finalized a month or more before that. We can consider data for the past years as *actuals*, but the numbers for 1995 and 1996 are perhaps better characterized as estimates. The arguments for including them in the database are, first for 1995, the end position for that year was reasonably well known several months into the year, when the data were prepared. For 1996, the February 1995 FYDP accompanied the President's Budget Submission for fiscal 1996. Hence the FYDP data were consistent with the detailed budget calculations affecting military personnel in that year. Nevertheless, it is reasonable to inquire as to the effects on the model's parameter estimates of first eliminating data for 1996 and then eliminating both 1995 and 1996.

The answer is that those changes had minimal effect on the estimates in the equation for officers. For enlistees, the estimates of β , the full-response parameter, increased by a substantial amount. Had the change been in the opposite direction, given that the original estimate (approximately 1.0) is quite high relative to earlier studies, that would have been cause for concern. But in light of this finding, the prudent course seems to be to accept the lower estimate.¹²

12. Another point to consider is that the estimate of β is obtained by forming the ratio of the estimates of γ_1 and $(1 - \gamma_2)$. Small changes in the estimate of γ_2 have a large effect on the estimate of β . Nevertheless, changes as large as those that occurred with the enlistee equation are a bit disquieting.

In [2], the model took the sum of officers and enlistees as the definition of afloat manning. We decided to assess the effects of that change on the results generated here. Since the number of enlistees afloat is some 11 times larger than the corresponding number of officers, the new definition had almost no effect on the enlistee equations. For officers, first in the simple regression, the total-military-afloat variable had less explanatory power than officers afloat (\bar{R}^2 of 0.536 as compared with the original 0.651). In the partial adjustment model, although the estimate of β fell from 0.941 to 0.084—owing to the 11-fold increase in the size of the at-sea variable—the elasticity estimate remained virtually unchanged. Because of these findings, and because we think the *a priori* argument is stronger for the original at-sea variables, we have elected to make no change to the earlier model specifications.

Concluding remarks

As noted at the outset, the main purpose of this analysis was to provide an empirical basis for estimating the indirect manning costs of different numbers and types of ships and aircraft. The resultant indirect cost factors—rounded to 0.9 for officers and 1.0 for enlistees—appear to rest on fairly solid footing, but they are very much larger than any that are known to be in use either now or in the past. As a representative case in point, the Naval Air Systems Command developed life-cycle cost estimates for several different aircraft and reported them in [9]. They estimated indirect personnel costs to be roughly 5 percent of direct manning costs. Although it may not be true of the particular application made of those estimates, differences this great could have an enormous effect on the outcome of a broad range of cost-related analyses.

The product of this work could have a number of applications beyond strict cost estimation. The statistical results documented in [2] have been used by Navy program analysts to size the shore establishment in response to force structure changes. Again, the magnitude of the response estimates in the present study greatly exceed those developed in that reference. Whether these new estimates will be institutionally acceptable is, at this time, simply unknown.

Appendix A: Time series manning data

Fiscal year	Officers afloat	Officers ashore	Enlistees afloat	Enlistees ashore
1980	19,758	43,300	222,608	236,961
1981	19,978	45,484	226,720	243,464
1982	20,472	46,801	234,239	246,947
1983	21,468	47,026	248,701	235,867
1984	22,011	46,845	248,267	243,021
1985	22,343	48,314	247,000	248,444
1986	22,775	49,276	250,798	253,591
1987	23,160	48,877	251,791	258,235
1988	23,689	48,738	249,300	266,026
1989	23,511	48,642	248,623	267,091
1990	22,315	50,773	244,774	260,195
1991	21,442	49,708	240,311	255,312
1992	20,364	48,748	211,127	257,251
1993	19,230	47,116	204,485	234,948
1994	17,251	44,499	186,782	215,844
1995	16,660	43,340	179,232	195,968
1996	15,441	43,364	166,378	198,817

Appendix B: Calculations supporting the sensitivity analysis

Tables 5 and 6 contain the estimates of parameter values and elasticities when the database spans 1980 through 1995 and 1980 through 1994, respectively.

Table 5. Estimates of the structural parameters and elasticities, with 1996 removed

Data set	α	β	λ	Elasticity
Officers	25,173	1.055	0.522	0.468
Enlistees	-202,871	1.891	0.210	1.783

Table 6. Estimates of the structural parameters and elasticities, with 1995 and 1996 removed

Data set	α	β	λ	Elasticity
Officers	27,258	0.968	0.535	0.433
Enlistees	-56,788	1.281	0.274	1.210

The following tables display the results of total military afloat for officers afloat and enlistees afloat in the first and second equations, respectively.

Table 7. Simple regressions of ashore manning on total-military-afloat manning

Data set	Intercept	Slope	S.E.E.	Dep. mean	Ind. mean	\bar{R}^2	D.W.
Officers	32,502 (9.75)	0.059 (4.41)	1,611	47,109	247,824	0.536	0.51
Enlistees	90,995 (4.04)	0.610 (6.77)	10,877	242,234	247,824	0.737	1.00

Table 8. Estimates of the partial adjustment model with total-military-afloat as the predictor variable

Data set	Est. of $\gamma_0 = \alpha\lambda$	Est. of $\gamma_1 = \beta\lambda$	Est. of $\gamma_2 = 1 - \lambda$	S.E.E.	Dep. mean	Ind. mean	\bar{R}^2	D.W.
Officers	12,232 (2.55)	0.038 (4.50)	0.545 (4.69)	862	47,347	248,165	0.849	1.91
Enlistees	6,242 (0.23)	0.339 (3.50)	0.622 (3.84)	7,986	242,564	248,165	0.866	2.08

Table 9. Estimates of the structural parameters and elasticities with total-military-afloat as the predictor variable

Data set	α	β	λ	Elasticity
Officers	26,884	0.084	0.455	0.440
Enlistees	16,513	0.897	0.378	0.918

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